TRADITIONAL EXPORT DEMAND RELATION: A COINTEGRATION AND PARAMETER CONSTANCY ANALYSIS

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Abstract

This study empirically estimates the critical parameters of the aggregate export demand function for Jordan by using annual time series data (1970-2004) and by applying both Johansen-Juselius and Saikkonen-Lütkepohl multivariate cointegration procedures. The empirical results confirm that there exists a unique and significant long-run equilibrium relationship among exports, foreign income, relative export price, and domestic output. Our estimation results show that income elasticity is much larger than unity while export price elasticity is slightly above one. The long-run estimate of the export price elasticity reveal that the Marshall-Lerner condition is satisfied for Jordan and currency devaluation may be effective in improving Jordanian exports and her trade balance. Moreover, domestic output has a positive and significant impact on Jordanian exports. Finally, tests for the parameter constancy suggested by Hansen and Johansen (1999) reveal that the hypothesis of stable long-run elasticities could not be rejected.

JEL Classification: F12, F13, C32

Keywords: Export demand, Elasticity, Cointegration, Stability analysis

I. Introduction

Jordan is classified as a lower-middle income country with a real per capita Gross Domestic Product in 2005 of $1,792. She is one of twenty one MENA countries that boost a population of 5.7 million, as of 2005.[1] It has an open economy that is based on few natural

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[1] MENA is Middle-East and North Africa countries.

resources such as potash and phosphate, with only 6% of its land considered arable, and water availability rank among the world's lowest. About half of her exports and a quarter of her imports are with her neighboring countries in the region. As such, much of its fortune is linked to those of the region, which has been beset by prolonged turmoil that adversely affected her economy.

With the help of the International Monetary fund (IMF), Jordan managed to stabilize her economy after the 1988-1989 economic and debt crisis. Main features of this crisis included, a sharp decline in her real GDP (-13% by 1989), mounting external debt (190% of GDP), current account deficit (14% of GDP), budget deficit (23% of GDP), sharp devaluation of the currency, the Dinar, and a severe depletion of her foreign exchange reserves (-73% decline in 1988). However, Jordan achieved major economic progress over the past decade. Her real GDP grew at about 4% during 2000-2005; inflation fell to low single digit levels; the public debt ratio was reduced substantially from 210% of GDP in 1990 to 83% of GDP by the end of 2005. This performance was accompanied by increased trade openness, export promotion policies and a surge of foreign direct investment.

Exports play an important role in the economic growth and development of many countries. In this respect, measuring income and price elasticties of export demand has received much attention because of the implications on trade policy and balance of payments (BOP) issues. The higher the foreign income elasticity of export demand, the more significant exports will be as an engine to economic growth. In addition, the higher the export price elasticity, the more competitive is the international market for exports of a particular country, and thus a real devaluation will be more successful in promoting export earnings. As such, an aggregate export demand estimate linking exports with a measure of foreign income and relative prices is important in many conventional trade models. Export demand elasticties are also important for meaningful export forecasts, planning, and policy formulation. Because of the importance of the trade sector for Jordan's economic growth and development and the ensuing implications on the balance of payments (BOP), the central aim of this paper is to estimate the long-run determinants of Jordan's exports during the period 1970-2004 (see Table 1 for some key trade indicators of Jordan).

Table 1. *Some key trade indicators for Jordan (1985-2004)*

Year	Exports as a % of GDP	Imports as a % of GDP	Imports as a % of C[a]	Trade Balance[b]	Foreign Exchange Reserves[c]	Foreign Exchange m.i.[d]
1985	40%	76%	66%	-1846	398.70	1.8
1986	28%	54%	51%	-1658	413.20	2.0
1987	33%	58%	57%	-1693	412.60	1.8
1988	43%	65%	66%	-1645	109.50	0.5
1989	56%	74%	78%	-793	459.70	2.6
1990	60%	90%	91%	-1272	847.80	3.9
1991	57%	80%	82%	-1277	824.70	3.8
1992	50%	82%	84%	-2002	750.20	2.7
1993	51%	82%	87%	-1982	1631.90	5.5
1994	49%	73%	82%	-1724	1691.90	6.0
1995	52%	73%	83%	-1702	1971.70	6.4
1996	53%	78%	82%	-2053	1758.50	4.9
1997	49%	72%	74%	-1822	2200.10	6.4
1998	45%	64%	66%	-1698	1749.60	5.5
1999	43%	61%	64%	-1468	2628.80	8.5
2000	42%	69%	66%	-2124	3330.60	8.7
2001	42%	67%	65%	-2062	3061.00	7.6
2002	45%	66%	65%	-1727	3975.00	9.4
2003	45%	67%	66%	-2027	5193.10	10.9
2004	48%	81%	79%	-3234	5264.80	7.7

Notes: [a] Imports as a % of Aggregate Consumption. [b] Trade Balance (millions of US dollars). [c] Foreign Exchange Reserves (year end) in millions of US dollars. [d] Foreign Exchange in months of Imports c.i.f *Sources are International Financial Statistics, Central Bank of Jordan and author's calculations.*

Moreover, unlike many studies for developed and developing countries that presumed (either explicitly or implicitly) stability of the estimated export demand function, this study formally tests the stability of the estimated parameters by applying Hansen-Johansen (1999) formal stability tests.

2. Recent empirical studies

A large volume of literature exists on the study of import and export demand functions for developing and developed countries. However, from the empirical literature surveyed, no recent study was found that estimates the determinants of the aggregate export demand function for Jordan. It is therefore only logical to briefly review most recent literature that is relevant to the theme chosen for this study (Johansen-Juselius cointegration framework).

Bahmani-Oskoee and Niroomand (1998) estimated trade elasticities for 29 developed and developing countries using Johansen's cointegration method. They found evidence of cointegration in 26 of them. Moreover, in most cases they found that the price elasticities are high and the sum of the absolute value of export and import demand price elasticities is greater than unity indicating that the Marshall-Lerner condition is satisfied.

Arize (2001) used different cointegration techniques including Johansen's to estimate the export demand function for Singapore, a newly industrializing economy. Their empirical results show evidence of a long-run and stable equilibrium relationship among exports and its determinants.

Guisan, M. and Cancelo, M. (2002) considered supply side determinants in addition to the traditional demand factors when estimating the determinants of exports for 25 OECD countries. Using data for the period 1960-1997, their econometric model included supply side variables such as domestic GDP, domestic private consumption and a measurement of human capital (proxied by educational levels of the population) in addition to foreign income and relative export price. Their results, in particular model 1, reveal that both external demand (foreign income), domestic GDP, and average years of schooling have a positive and significant impact on exports, while both consumption and relative prices have a significant and negative impact on exports.

Khedhiri and Bouazizi (2007) estimated demand elasticities for Tunisian exports to the major European trading partners. Using panel cointegration technique, the authors found a significant relationship between real exchange rate index, foreign income and Tunisian exports. With respect to Tunisian exports, Khedhiri and Bouazizi found foreign income to be elastic (2.95), while real exchange rate to be inelastic (-0.162). They conclude that devaluation of Tunisian currency will slightly improve export demand but may be very costly since it will have a depressing effect on domestic output.

This study contributes to the existing literature on trade elasticities in five ways. First, it applies recent techniques in time series analysis that include, among others, using Johansen- Juselius (1990) and Saikkonen-Lütkepohl (2000abc) multivariate cointegration framework. Second, this study tackles the issues of unit root tests, choice of lag order, and deterministic components that are critical to cointegration tests such as Johansen and Juselius. Third, this study uses most up to date data and relatively a large sample size to estimate the long–run relations between exports and its determinants. As such, in this study, the focus is on the long-run relations of the aggregate export demand functions of Jordan, mainly considering policy implication issues. Fourth, the econometric model we use to estimate the Jordanian exports, takes into account supply side determinants in addition to the traditional ones, i.e., level of external demand and export relative price, by including the level of domestic GDP. Strong economic performance measured by domestic GDP will have a direct and positive effect on exports (a hypothesis that is maintained by proponents of growth-led exports), but will also indirectly have a positive impact on the level of exports through levels of education.[2] Fifth, unlike many previous studies for developed and developing countries that assumed stability of the cointegrating vector, this study formally tests the constancy of the estimates by applying formal stability tests suggested by Hansen-Johansen (1999).

3. The model and the methods

To estimate the rest of the world’s demand for Jordanian exports, the traditional long-run specification of export demand function that relates the volume of a country’s exports to world buying power (foreign income) and the ratio of the price of its exports to the world export price will be included. Furthermore, following Guisan, M. and

[2] The author would like to acknowledge the editor of this Journal. She correctly pointed out the importance of including supply side factors such as domestic GDP when estimating the demand for exports.

Cancelo, M. (2002), domestic GDP proxied by the industrial production index will be included to capture supply side effects on Jordanian exports. The aggregate export demand function is specified as follows:

$$\ln X_t^d = \gamma_0 + \gamma_1 \ln YW_t + \gamma_2 \ln(\frac{PX}{PXW}) + \gamma_3 \ln YD + \varepsilon_t \qquad (1)$$

where X^d is the volume of Jordan's export, YW is the rest of the world income, PX is Jordan's export price, PXW is world export price, YD is domestic GDP and ε_t is a serially uncorrelated random term, and ln stands for the natural logarithm of the relevant variables. We expect γ_1 to be positive, an indication that as world income rises, their demand for goods and services increases, including those of Jordan, γ_2 to be negative reflecting the fact that as Jordan export price rises relative to world export price, Jordanian goods and services become more expensive to foreign buyers. Moreover, we expect γ_3 to be positive, an indication that strong economic performance measured by higher GDP will directly lead to higher exports.[3]

The loglinear form is chosen, since it is found to be the most appropriate function form for both import and the export demand functions in many empirical studies (Khan and Ross 1977, Boylan et al 1980, Emran and Shilpi 1996). It also has the added advantage of reducing hetroskedasticity (Maddala 1992).

4. The empirical framework

Equation 1 can be estimated by the standard regression method if all variables in the equation are stationary and the errors have a zero mean and finite variance. However, in the presence of nonstationarity (integrated variables), there might occur what Granger and Newbold (1974) called a spurious regression, hence, the usual *t* and F tests may give misleading results (Engle and Granger 1987). Thus, the first step in this study is to investigate the integration properties of the time-series data. If the variables are

[3] If the increase in world income is due to an increase in their production of import substitute goods, γ_1 could be negative. Moreover, strong economic performance measured by higher domestic income (GDP) can also indirectly lead to higher exports through improved levels of education.

found to be integrated (nonstationary), then the issue is to what degree they are integrated.[4] If the variables in the data are integrated of order one, I(1), we proceed to test whether they are cointegrated using Johansen and Juselius (1990, 1992) cointegration framework to estimate the export demand function in equation 1.

Unit-root testing

The Phillips-Perron (1988) unit root test is used in this study, in conjunction with the Augmented Dickey-Fuller (1979) test to address the issue of integration of the time-series data.[5] The Phillips and Perron (*PP*) is considered here because it accounts for possible correlation in the first differences of the time-series using a nonparametric correction, and allows for the presence of a non-zero mean and a deterministic time trend. In addition, Perron (1989) has suggested that Augmented Dickey-Fuller (*ADF*) tests may falsely conclude the presence of a unit root in a time-series subject to a structural break. In the Annex we analyze some features of this test.

Cointegration tests: the Johansen-Juselius (JJ) method

If the variables in equation 1 are integrated (nonstationary), the next step is to test for cointegration properties. In general, a set of variables are cointegrated if a linear combination of the integrated series is stationary, i.e., if $Y_t \sim I(d)$ and $X_t \sim I(d)$, the following regression is run:

$$Y_t = \beta X_t + \varepsilon_t \tag{3}$$

if the residuals, ε_t, are $I(0)$, then X_t and Y_t are said to be cointegrated. Johansen's approach is used since it allows for the presence of multiple cointegration relationships, *r*, in a single-step procedure to be estimated and tested for. In the Annex we analyze this test. We consider Saikkonen and Lütkepohl (S&L) tests, in conjunction with Johansen, to account for the possibility of any structural breaks (shifts) in the data generating process (DGP) that is

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[4] If a time series requires differencing d times before it becomes stationary, it is integrated of order d, i.e., I(d).

[5] The joint use of both tests attempts to overcome the common criticism that unit root tests have limited power in finite samples to reject the null hypothesis of nonstationarity.

due to internal or external shocks that the Jordanian economy may have experienced during the time of the study.[6]

4. Data and variables

In estimating the long-run parameters of the export demand function for Jordan, this study employs the latest econometric techniques, i.e., Johansen-Juselius (1990) and Saikkonen-Lütkepohl (2000abc) multivariate cointegration tests and uses recent available annual data (1970-2004). For estimating the export demand function in equation 1, the variables used are as follows: *XV* is Jordan's index of volume of exports, YW is world income proxied by the index of industrial production in industrial countries, PX is export price proxied by export price index; PXW is world export price proxied by world export price index, PXPXW is export relative price (PX divided by PXW), and YD is domestic GDP proxied by the industrial production index. All data used in this study were obtained from various issues of the international financial statistics, published by the International Monetary Fund (IMF) and the yearly statistical series (1964-2003) of the Central Bank of Jordan, all are expressed in natural logarithm and the base year is 2000 (2000=100).

5. Empirical results

Test results for unit roots

Table 2, in the Annex, provides the unit root test results of the Phillips-Perron (PP) and the Augmented Dickey-Fuller (ADF) tests for the export demand variables. The test statistics τ_μ and $Z(t_{b1})$ (intercept or a drift) and τ_τ and $Z(t_b)$ (both a drift and a linear time trend) are reported in Table 2. It can be seen that in none of the level variables are the computed *PP* and *ADF* statistics less than its 95% critical value (except for *LYD* for drift case only). Therefore, all

[6] Jordan experienced an economic crisis late 1980s that led to economic and structural reforms through IMF-supported adjustment programs (six in total) during the period 1989-2004. Main features of the crisis were, sharp decline of GDP (-13% by 1989) and double-digit inflation rate (25%). Therefore, we include a shift dummy, D1987, in the cointegration test proposed by Saikkonen and Lütkepohl and the VECM estimations (results are not sensitive to the inclusion of D1987).

variables are nonstationary at the 95% level of significance. In first difference, the calculated *ADF* and *PP* for the same variables are less than the 95% critical values. We can conclude that all first differenced variables are stationary or *I(0)*. Thus all level variables are integrated of the same order or *I(1)*. This conclusion is confirmed when implementing the Johansen (1995) multivariate stationarity test (see Table 6).

Selection of optimal lag order and deterministic components

Because of the sensitivity of LR tests to both the presence of deterministic components (a constant, a trend, seasonal and other dummies) and to the choice of the appropriate lag order, *p*, of the VAR model, one needs to specify the appropriate lag order before discussing the characteristics of the Π matrix that contains the cointegrating vector or the long-run relations. It is been proven that working with a low-order model that does not capture the serial dependence in the data may lead to size distortions, whereas choosing an unnecessarily large lag-order may spoil the power of the test.

Lütkepohl (2004) among others suggest using model selection criteria such as Akaike information criterion (AIC) and Schwarz information criterion (SC). The general approach to using the criterion is to fit VAR (m) models with orders $m = 0, \ldots, p_{\max}$ and to choose an estimator of the order p that minimizes the “preferred” criterion. Taking into account the data frequency and the number of available observations, we consider a maximum lag order of four. The VAR lag order selection criteria we followed show that HQ (Hannan-Quinn information criterion) and FPE (Final prediction error) favor a VAR specification with three level lags. SC (Schwarz information criterion) favors a one level lag, while AIC (Akaike information criterion) favors a four level lag; therefore we apply the cointegration test for the export demand function using specification suggested by HQ and FPE.[7]

[7] Choosing the order too small can lead to size distortions for the tests while selecting too large an order may imply reductions in power (Lütkepohl 2004). The lag length, *P*, in the VAR corresponds to a lag length of ***P*-*1*** in the VECM.

The decision regarding the deterministic components in the model is not easy to determine in advance. An important feature of the cointegrated VAR in equation 6 is that it includes both level and differenced variables. The asymptotic distribution of the test for cointegration depends on the assumptions made regarding deterministic components in a model. Johansen (1992) suggests testing the joint hypothesis of both the rank order and the deterministic components based on what so called the Pantula principle. That is, all models that can realistically be considered (known in the literature as models 2, 3, and 4) are estimated and the results are presented from the most restrictive model (r = 0 and model 2) to the least restrictive alternative ($r = k$ -1 and model 4).[8] The test procedure then is to move from the most to the less restrictive model and at each stage to compare the trace test statistic to its critical value and only to stop on the first occasion the null hypothesis is not rejected.

Determination of the rank of Π

Tables 3 and 4, in the Annex, present Johansen and Saikkonen-Lütkepohl (S&L) cointegration results for the export demand variables, i.e., *LXV*, *LYW*, *LPXPXW*, and *LYD*. The results show that both Johansen and S&L test statistics reject the hypothesis of no cointegration at 5% level of significance. Furthermore, the two tests for cointegration suggest a unique cointegration relation among real exports (*LXV*), foreign income (*LYW*), exports relative price (*LPXPXW*), and domestic output (*LYD*). Hence, in Jordan a long-run relationship exists between exports and its determinants.

Estimated long-run relations

The existence of one cointegration vector implies that an economic interpretation of the long-run export demand function can be made by normalizing the estimates of the unconstrained cointegrating vector on exports. Table 5 reports the coefficients of this vector. Moreover, the table reports the test for the exclusion of variables

[8] The 3 models that can realistically be considered are model 2 , which includes intercepts in the cointegrating vector and no linear trend in the level of the data; model 3, which includes deterministic trends in levels (no trend or intercept in cointegrating relation) and model 4, which allows for trends in cointegrating relations.

from the cointegration vector matrix. After all, the observed cointegration among the variables could be due to a strong relationship among some of the variables but not all. The exclusion test is performed by placing zero restrictions on the parameters and the test statistic used is the likelihood ratio (LR) test suggested by Johansen-Juselius (1990). If a variable can be excluded from the cointegration space, it implies that it evolves independently and as such not integrated with the other variables in the system.

The estimates of the long-run parameters for the export demand system show that the null hypothesis of long-run zero restrictions is rejected for all variables at 5% significance level. These LR test results suggest that all explanatory variables are significant and can be included in the export demand function for Jordan. Moreover, VAR diagnostic tests such as the LM test for autocorrelation, normality test, and multivariate ARCH-LM test are performed and reported. All diagnostic tests show no evidence of serial correlation, non normality or any problem of heteroskedasticity. Moreover, Johansen (1995) multivariate unit root test rejects the null hypothesis of stationarity for *LXV*, *LYW*, *LPXPXW*, and *LYD*.

Table 5. ML estimates of the cointegrating vector (Normalized on LXV)

Variable	Cointegrating Vector (β)	S.E.	[t-stat]	LR Exclusion Test	(Prob.)	Johansen Stationarity test	(prob.)
LXV	1	-		.76	(0.00)	19.32	(0.00)
LYW	-2.15	0.255	[-8.45]	8.42	(0.00)	21.11	(0.00)
LPXPXW	1.031	0.139	[7.393]	12.6	(0.00)	8.47	(0.03)
LYD	-0.54	0.070	[-7.71]	5.11	(0.02)	19.04	(0.00)
C	2.73	0.791	[3.455]	2.47	(0.11)	-	

Nonormality Test-Jarque-Bera (p-value)	*0.98*
LM Serial Correlation (Lag 1, p-value)	*0.03*
LM Serial Correlation (Lag 12, p-value)	*0.46*
Multivariate ARCH-LM Test	*0.28*

Notes: Figures in brackets are probability values showing the exact level of significance.

Based on the estimated coefficients reported in Table 5, the long-run equilibrium export demand function is given by the following equation (written in explicit form):

$$LX = -2.73 + 2.15LYW - 1.03LPXPXW + 0.54LYD \qquad (6)$$

In equation 6, world income, relative price, and domestic income elasticities carry their theoretical expected signs and are significantly different from zero. The elasticity coefficient of foreign income (*LYW*) is positive and is significantly different from zero. This implies that as world income grows, Jordan will capture a larger portion of world exports. For instance, a 1 percent increase in foreign income induces a 2.15 percent increase in export earnings, all else unchanged. In addition, this high and significant world income elasticity may be a reflection of the greater variety and quality of Jordanian exports. The estimated relative price elasticity is negative, significant, and is near unity (-1.03), which means that changes in relative export price, ceteris paribus, will lead to nearly no change in export earnings. Like many NIEs (newly Industrialized Economies), Jordan could increase her share of world exports without experiencing drastic fall in her terms of trade. In this study, world income elasticity is significant and much larger than unity, thus, external demand is vital to Jordan's export growth and the near unitary export price elasticity confirms that Jordanian exports are demand-constrained. Finally, the elasticity of domestic GDP with respect to exports is positive, much larger than unity, and is significantly different from zero; an indication that supply side determinants play a significant role in Jordan's export growth. The positive and significant coefficient of domestic GDP lends support to the GLE (growth-led exports) hypothesis. Strong economic performance in Jordan leads to higher exports and possibly improvements in Jordan's trade balance. Comparing this study's income elasticity of exports for Jordan with those obtained for other developing countries by Bahmani-Oskoee and Kara (2005), Jordan's income elasticity of exports is much higher than the income elasticities for all developing countries in their study. This is important to Jordan since high-income elasticity of exports implies that as world income grows, Jordan will be in a position to capture a larger percentage of world exports, thus narrowing the balance of payments gap. One case similar to Jordan in Bahmani-Oskoee and Kara study is that of South Korea, which has a high-income elasticity of three. Jordan, like South Korea, with export promotion policies as its engine towards growth, will benefit from this high-income elasticity. Finally, this high-income elasticity of exports lends support to the export promotion and trade liberalization policies that Jordan enacted in the last decade.[9] It is also important to point out that even in the absence of estimated price elasticity for Jordan's imports; the Marshall-Lerner (M-L) condition is satisfied, since the export price elasticity is 1.03.[10] This in turn supports the case for devaluation to improve Jordan's trade balance.

7. Stability analysis

In contrast to most if not all previous studies that estimated long-run and short run trade elasticities for developing and developed countries, this study applies formal tests to investigate the parameter constancy issue. Having estimated the cointegrating vector, it is necessary to test whether the estimated long-run "elasticities" are stable. Long-run stability means that those parameters of the cointegration relationship are invariant overtime. As has been emphasized by Brüggemann et al. (2003), it is of some importance to formally investigate the stability of the cointegrating vectors further, once a long-run relationship has been identified. For cointegrated VAR models, Hansen and Johansen (1999) suggested applying a fluctuation test to the nonzero eigenvalues of the reduced rank matrix. The fluctuation test rejects stability when the recursively estimated eigenvalues fluctuate excessively. The test may be applied to the eigenvalues themselves, λ_i, giving rise to the test statistic Sup λ_i, or to the transformation $\xi_i = \log(\lambda_i/(1-\lambda_i))$, giving rise to the test statistic Sup ξ_i. It is also possible to evaluate the constancy of the

[9] Husein (2008) tested the export-led growth (ELG) hypothesis for Jordan and found evidence of bidirectional causality between exports and economic growth.

[10] If the sum of the absolute value of export and import demand price elasticities is greater than unity, devaluation could improve a nation's trade balance

eigenvalues jointly by considering the sum of the transformed eigenvalues, which gives rise to the test statistic $Sup\sum_{i=1}^{r}\xi_i$.

To examine the constancy of the cointegration space, we consider two types of Nyblom tests studied by Hansen and Johansen (1999), $SupQ_s$ and $MeanQ_s$. The first (supermum) test statistic, is based on the maximum value of a weighted LM-type test statistic over the experimentation period, and the second (mean) test, on the average of this statistic.[11] Finally, we consider the constancy of the (unrestricted) Φ, Γ_1, and α parameters in equation 6 using the fluctuation test due to Ploberger-Krämer-Kontrus (1989). It is worth mentioning that all formal tests do not require trimming of the sample, however, we use 30% of the sample as a base period and examine constancy over the remainder.When testing stability of a subset of parameters, an important issue is how to treat the remaining parameters in the model. One approach is to fix the latter parameters at the full sample estimates, and the other is to update them along with the parameters being analyzed.[12]

Constancy of the non-zero eigenvalues

Panel A of Table 7, in the Annex, report the Hansen-Johansen fluctuation tests, Sup λ_i and Sup ξ_i, conditional on the full sample estimates of the deterministic and lagged parameters, Φ and the Γ_1.

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[11] The LM-type statistic for the two Nyblom tests can be calculated using two different methods. The first suggested by Hansen-Johansen (1999) and involves a first order Taylor expansion of the score function; the other, suggested by Brüggemann et al. (2003), uses the scores directly. Brüggemann et al. (2003) and Warne (2005) suggest that the score version to computing the LM-type statistic is superior to that of Hansen-Johansen (*HJ*) since the latter suffer from numerical problems in simulation exercises, leading to small sample distributions that are far away from the limit distributions.

[12] According to Brüggemann et al. (2003), updating parameters of interest is ideally the preferred approach, but given the finite samples and that the tests are based on asymptotic theory, the more parameters we update the more likely is that the asymptotic distributions provide poor approximations of the unknown small sample distributions. Therefore, the analysis is extended with bootstrapped empirical distributions for all stability tests.

As can be seen, the null hypothesis of constant eigenvalue cannot be rejected using both asymptotic and bootstrapped critical values over the test period for Sup λ_i, and using bootstrapped critical value for Sup ξ_i test at 1% level. Moreover, constancy of the eigenvalue cannot be rejected if instead the Φ and the Γ_1 are updated as seen in Panel B of the same table for Sup λ_i and Sup ξ_i tests using bootstrapped critical values. In either case, it can be concluded that the non-zero eigenvalue is constant for the cointegrated VAR in equation 6.

Constancy of the cointegration space

Table 8 panels A and B, in the Annex, reports the two Nyblom type tests, $SupQ_s$ and $MeanQ_s$, for testing the stability of the cointegrating vector, β. Panel A tests the constancy of β conditional on full sample estimates of Φ and Γ_1. As can be seen, the hypothesis of inconstant β is strongly rejected for the cointegrated VAR in equation 6, using both the asymptotic and bootstrapped critical values. Both the Nyblom supermum ($SupQ_s$) and mean ($MeanQ_s$) tests are far below their five percent asymptotic and bootstrapped critical values of 2.71 and 0.86, respectively. Similarly, Panel B of Table 8 reports the results of the $SupQ_s$ and $MeanQ_s$ when the Φ and Γ_1 parameters are updated. Both tests fail to reject the null hypothesis of constant β using both the asymptotic and bootstrapped critical values. Based on the two Nyblom type tests, we conclude that the cointegration space for the cointegrated VAR in equation 6 is constant for the test period.[13]

Table 9: *Ploberger-Krämer-Kontrus fluctuation tests for the constancy of Φ, Γ1, and α for the cointegrated VAR with 3 lags and one cointegration relation.*

Equation	S(10)	asymptotic *p*-value	bootstrap *p*-value
LXV	2.38	0.00	0.69
LYW	2.09	0.00	0.78
LPXPXW	2.39	0.00	0.68
LYD	2.50	0.00	0.65

[13] For more detailed analysis on all the tests, see Hansen-Johansen (1999), Brüggemann et al. (2003), and Lütkepohl, H. and Krätzig, M. (2004) and Lütkepohl, H (2005).

Furthermore, Table 9 reports fluctuation tests for the constancy of (unrestricted) Φ, Γ_1, and α due to Ploberger-Krämer-Kontrus (1989) in equation 6. Based on bootstrapped p-values, there are no signs of non-constancy in *LXV*, *LPXPXW*, and *LYW* equations. Based on the stability analysis reported in Tables 7, 8 and 9, the eigenvalue fluctuation, Nyblom-type, and Ploberger-Krämer-Kontrus tests for the cointegrated VAR model, the empirical results strongly rejects instability of the eigenvalues, the cointegrating relationships, and Φ, Γ_1, and α parameters.[14]

8. Concluding remarks

The primary objective of this study has been to estimate the critical parameters of the export demand function for Jordan. The empirical results obtained show that real exports, relative price, real foreign income, and domestic GDP are cointegrated and that the cointegrating vector is unique. The results also show that foreign income is a significant variable in explaining the demand for exports and that foreign income elasticity is much larger than unity, an indication that an increase in world's income will more than proportionately increase Jordanian exports. Export demand for Jordanian goods is unitary price elastic. Other things being equal, a change in export price leads to nearly no change in export earnings. An indication that that the status of external demand plays an important role in determining Jordanian exports. Given the large income elasticity, an increase in quality and variety of her exports will lead to a considerable increase in her share of the world market.

Policies that improve the non-price competiveness of Jordanian exports and stimulates better understanding of her external demand needs to be implemented. Moreover, in this study, supply side determinant, i.e., domestic GDP proxied by industrial production index has a positive and significant impact on exports. An indication that domestic production, i.e., strong economic performance, leads to increased exports, which lends support to the GLE hypothesis. Finally, Hansen-Johansen formal stability tests indicate that the

[14] For more detailed analysis on constancy tests, see Hansen-Johansen (1999), Brüggemann et al. (2003), and Lütkepohl, H. (2004) and (2005).

estimated relative export price, foreign income, and domestic GDP elasticities are stable over time; this is essential since they are used as a tool for policy formulation. Finally, the Marshall-Lerner condition for Jordan seem to be satisfied since export price elasticity is slightly larger than unity, hence, devaluation could improve the Jordanian trade balance through reductions in her imports and expansions in her exports. However, as Arize (2004) points out, a small and frequent devaluation approach is likely to be a reasonable way of avoiding any disruptive effects associated with large and frequent devaluations. Finally, devaluations may be accompanied with other appropriate policy instruments.

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Annex

Phillips-Perron test

In general, the *PP* test is based on the estimate of the following regression(s):

$$\Delta Y_t = a_0 + a_1 t + \alpha Y_{t-1} + \varepsilon_t \qquad (2)$$

where *Y* is a time series, a_0 is an intercept or drift term and *t* represents a linear time trend, and Δ is the difference operator. Equation 2 includes both a drift and a linear time trend (becomes a random walk with a drift when a_1 is zero). Once the regression is estimated, the null hypothesis of unit root may be tested:

$$H_0 : \alpha = 0 \qquad\qquad H_A : \alpha < 0$$

while the ADF test corrects for higher order serial correlation by adding lagged differenced terms of Y_t on the right hand side of equation 2, the PP test makes a correction to the *t*-statistic of the α coefficient by using heteroskedasticity autocorrelation consistent estimates:

$$\gamma_j = \frac{1}{T}\sum_{t=j+1}^{T} \hat{\varepsilon}_t \hat{\varepsilon}_{t-j} \qquad\qquad w^2 = \gamma_0 + 2\sum_{j=1}^{p}\left(1 - \frac{j}{p+1}\right)\gamma_j$$

where $\hat{\varepsilon}_t$'s are the estimated residuals from equation 2, *T* is number of observations, and *P* is a proper truncation lag that assures white noise residuals. Finally, the above calculated estimates are used to compute the PP t-statistic, t_{pp}:

$$t_{pp} = \frac{\gamma^{1/2} t_b}{w} - \frac{\left(w^2 - \gamma_0\right) T s_b}{2 w s}$$

Where *s* is the standard error of the test regression in equation 2, S_b and t_b are the estimated standard error of α and the standard *t*-statistic for testing the null that $\gamma_1 = 0$, respectively. The asymptotic distribution of the *PP t*-statistic is the same as the ADF *t*-statistic, hence, the same critical values of both tests are used to determine if the null hypothesis is rejected or not. If the calculated *PP* or *ADF* test statistic for Y_t is less than its critical value, then the series Y_t is said to be stationary or integrated of order zero, *I(0)*. If that is not the case, then the tests are performed on the Y_t's first differences. If the first differences are found to be stationary, then the order of integration is said to be one, *I(1)*.

Cointegration tests: Johansen-Juselius and Saikkonen-Lütkepohl

Consider the following basic vector auto regressive model of order *p*, VAR (*p*):

$$Y_t = A_1 Y_{t-1} + ... + A_p Y_{t-p} + \mu_t \qquad (4)$$

where Y_t is a vector of nonstationary variables, A_i's are ($K \times K$) matrix of parameters, and $\varepsilon_t = (\varepsilon_{1t}, \ldots, \varepsilon_{kt})$` is an unobservable error term assumed to be i.i.d. (independently and identically distributed) *k* dimensional Gaussian term with $\varepsilon_t \sim (0, \Sigma_\mu)$. Since Y_t is *I(1)*, the VAR can be written in the first-differenced error-correction form by subtracting Y_{t-1} from both sides of equation 2 and rearranging terms such as:

$$\Delta Y_t = \Gamma_1 \Delta Y_{t-1} + ... + \Gamma_{p-1} \Delta Y_{t-p+1} + \Pi Y_{t-1} + \mu_t \qquad (5)$$

where $\Gamma_i = - (A_{i+1} + \ldots + A_p)$ and $\Pi = - (I_k - A_1 - \ldots - A_p)$ for i = 1, …, *p* - 1

because ΔY_t is I(0), it follows that ΠY_{t-1} is the only term that includes I(1) variables. Hence, ΠY_{t-1} must be I(0) and as a result it contains the cointegration relations. The focus of the Johansen and Juselius technique is on the parameter matrix Π, which contains information about the long-run relationship among the variables in the data vector.

The rank of this matrix Π, rk(Π), determines the number of cointegrating vectors in the VAR system. If matrix Π has a full rank, i.e., rk(Π)= *k*, the vector Y_t is stationary. Instead, if matrix Π has a rank that equals zero, rk(Π)=0, then Π is a null matrix and equation 5 corresponds to a traditional VAR model in first difference. Finally, if matrix Π has a reduced rank ($0 < r < K$), then there exists $k \times r$ matrices α and β, each with rank r such that $\Pi = \alpha \beta`$ and $\beta`Y_t$ are *I(0)* even though Y_t itself is *I(1)*. r is the number of cointegrating relations and each column of β is the cointegrating vector. The matrix α contains the weights attached to the cointegrating relations in the individual equations of the model and sometimes is referred to as the loading matrix. In this case, equation 5 is a vector error-correction model of order p - 1, VECM (p - 1).

Several extensions of the VECM in equation 5 are usually necessary to represent the main characteristics of a data set of interest. Including deterministic terms such as a constant, a linear trend term, and seasonal and other dummy variables, may be required for a proper representation of the process. A general VECM that includes all such terms is:

$$\Delta Y_t = \Gamma_1 \Delta Y_{t-1} + \ldots + \Gamma_{p-1} \Delta Y_{t-p+1} + \alpha \beta' Y_{t-1} + \Phi D_t + \mu_t \tag{6}$$

where D_t contains all regressors associated with deterministic terms and Φ is a matrix of parameters.

The Johansen-Juselius and Saikkonen-Lütkepohl (S&L) cointegration techniques allow estimation of the cointegrating relationships among the integrated variables using a maximum likelihood (ML) procedure that tests for the rank of Π and estimates the parameters of β. Both cointegration approaches allow for the presence of multiple cointegration relationships, r, in a single-step procedure to be estimated and tested for. A general Vector Error Correction Model (VECM) that centers on the cointegration relation is of the form The cointegrating rank, r, can be tested using a likelihood ratio (LR) tests that is known as the trace test. The LR test, (λ_{Trace}), for the null hypothesis that there are at most r cointegrating vectors is computed as follows:

$$\lambda_{Trace}(r) = -T \sum_{i=r+1}^{k} \ln(1 - \hat{\lambda}_i)$$

where $\hat{\lambda}_{r+1}, \ldots, \hat{\lambda}_k$ are the $k - r$ smallest estimated values of the characteristic roots (eigenvalues) obtained from the estimated Π matrix.

It has been suggested that the above two tests are sensitive to both the presence of deterministic components (a constant, a trend, seasonal and other dummies) and to the choice of the appropriate lag order, *p*, of the VAR model. As such, one needs to specify the appropriate lag order, *P*, of the VAR representation of the cointegration equation such that the residuals are uncorrelated and homoscedastic.

Table 2. *Results of PP and ADF unit root tests for the export demand variables*

	Phillips-Perron Test				ADF Test			
	Level		1st-difference		Level		1st-difference	
	$Z(t_{b1})$	$Z(t_b)$	$Z(t_{b1})$	$Z(t_b)$	τ_μ	τ_τ	τ_μ	τ_τ
LXV	-1.00	-1.68	-6.72	-7.00	-.095	-1.68	-6.80	-7.00
LYW	-2.15	-2.84	-6.82	-9.47	-0.54	-3.13	-4.15	-4.08
LPXPXW	-1.98	-2.13	-6.09	-5.95	-1.98	-2.13	-6.09	-5.99
LYD	-4.10	-1.90	-4.90	-5.94	-3.90	-1.91	-2.94	-5.93

Note: The 1% and 5% McKinnon (1993) critical values for rejecting the null hypothesis of unit root for both ADF and PP are -3.63 and -2.95 for random walk with drift (τ_μ, $Z(t_{b1})$) and -4.25 and -3.54 for drift and linear trend (τ_τ, $Z(t_b)$). Optimal lag orders chosen by Akaike information criterion (AIC) for the *ADF* test and by Newey-West automatic truncation lag for the *PP* test.

Table 3. *Johansen-Juselius λ_{Trace} ML Cointegration tests*

		Model 2		**Model 3**		**Model 4**	
H_O:	H_A:	λ_{trace}	CV	λ_{trace}	CV	λ_{trace}	CV
$r = 0$	$r \geq 1$	68.34	53.4	53.2	46.6	90.48	62.7
$r \leq 1$	$r \geq 2$	31.96	33.9	18.3	29.3	48.41	42.9
$r \leq 2$	$r \geq 3$	8.24	19.9	6.10	14.8	13.96	28.3
$r \leq 3$	$r \geq 4$	2.58	9.1	2.04	2.52	2.16	10.6

Notes: r = cointegration rank. The 5% critical values of Johansen cointegration tests are obtained by computing the respective response surface according to Doornik (1998).

Table 4. *Saikkonen and Lütkepohl (S&L) Cointegration Tests*

	Model 2	**Model 3**	**Model 4**

H_O:	H_A:	LR	CV	LR	CV	LR	CV
$r = 0$	$r \geq 1$	51.25	40.1	50.85	35.7	49.65	45.3
$r \leq 1$	$r \geq 2$	23.84	24.1	28.23	20.9	21.04	28.5
$r \leq 2$	$r \geq 3$	4.53	12.2	8.85	9.84	14.25	15.7
$r \leq 3$	$r \geq 4$	1.14	4.1	-	-	2.65	6.79

Note: r = cointegration rank or the number of cointegrating vectors. (2) Shift dummy, D1987, is included in the deterministic component of models 2 and 4. Critical values (5%) are according to Trenkler, C. (2004).

Table 7: *Hansen-Johansen fluctuation tests of the stability of the non-zero eigenvalue for the cointegrated VAR with p = 1 and one cointegration relation over the period 1987-2004*

(A) Conditional on $\hat{\Phi}^{(T)}$ *and* $\hat{\Gamma}_1^{(T)}$

bootstrap CV 5%	bootstrap CV 1%	asymptotic CV 5%	asymptotic CV 1%	Sup λ_i
1.29	1.25	1.63	1.35	1.61

Sup ξ_i				
2.09	1.95	2.78	1.35	1.61

(B) Updating of $\hat{\Phi}^{(T)}$ *and* $\hat{\Gamma}_1^{(T)}$

Sup λ_i			
1.81		2.14	2.93
	1.35	1.61	

Sup ξ_i			
30.9		35.5	48.5
	1.35	1.61	

Table 8: *Nyblom supermum test and Nyblom mean test for the constancy of long-run parameters, β, over the period 1987-2004*

(A) Conditional on $\hat{\Phi}^{(T)}$ *and* $\hat{\Gamma}_1^{(T)}$

asymp boot		asymp	bootstrap $SupQ_s$		5% *CV*
	5% *CV*		$MeanQ_s$	5% *CV*	5% *CV*
1.26			2.71	2.03	
	0.490	0.86		0.92	

(B) Updating of $\hat{\Phi}^{(T)}$ *and* $\hat{\Gamma}_1^{(T)}$

$SupQ_s$					
1.68			2.58	3.03	
	0.717	0.87		1.36	

Obs.	XV	PX	PXW	YD	YW	PXPXW
1970	7.984750	28.25430	25.88540	10.15530	49.11590	109.1515
1971	7.280560	26.28570	27.12240	12.04020	50.01680	96.91510
1972	9.056990	31.66830	29.67030	14.73170	53.22910	106.7340
1973	9.833830	34.72470	36.51080	16.18690	58.24220	95.10802
1974	13.95200	72.02890	50.58390	16.88140	58.17250	142.3949
1975	11.98070	82.25110	54.09710	18.10360	53.40490	152.0435
1976	14.49070	73.71260	54.81170	22.62950	57.50070	134.4833
1977	17.47950	74.28480	59.95440	23.49320	60.09270	123.9022
1978	18.80660	77.07300	65.71300	28.84820	62.35320	117.2873
1979	23.24100	79.00780	78.31380	34.54880	65.07710	100.8862
1980	29.01170	94.63590	94.37930	41.23690	64.87160	100.2719
1981	33.82490	97.88520	94.80560	48.08330	64.67680	103.2483
1982	34.61510	100.2550	91.04210	49.68690	63.02330	110.1194
1983	34.35480	88.94670	86.40060	52.14850	64.21800	102.9469
1984	48.21720	91.54660	84.66260	59.75500	68.55810	108.1311
1985	50.22630	86.50790	82.93950	64.00450	70.18390	104.3024
1986	51.53220	83.98810	89.42680	64.87400	71.07050	93.91827
1987	61.22590	80.59000	98.45650	70.90560	73.65300	81.85341
1988	68.60910	84.91390	103.7880	65.13600	77.87870	81.81476
1989	72.32590	86.32940	105.6170	68.35480	80.49760	81.73817
1990	69.61360	88.98300	112.6120	68.78090	82.29740	79.01733
1991	61.17560	96.51830	110.5020	67.96610	82.21500	87.34530
1992	67.20280	93.15180	112.7300	73.26930	81.42520	82.63266
1993	72.57700	92.33580	107.0360	79.02480	80.52790	86.26612
1994	77.70010	95.90250	110.1370	83.44920	83.74020	87.07564
1995	84.46000	111.3730	120.9110	93.79690	85.89600	92.11155
1996	82.20670	117.0130	118.7450	88.95690	87.73230	98.54141
1997	87.80110	113.6110	111.3990	92.10710	91.95100	101.9857
1998	90.28750	107.2780	105.3010	93.97080	93.06940	101.8775
1999	92.85160	104.6310	103.6740	96.23090	95.39810	100.9231
2000	100.0000	100.0000	100.0000	100.0000	100.0000	100.0000
2001	123.2910	101.3570	96.19140	106.6090	96.87530	105.3701
2002	141.8030	101.7010	96.91700	118.4880	96.43100	104.9362
2003	152.3700	101.8900	106.8070	108.4610	97.50730	95.39637
2004	189.6660	113.7050	116.3640	121.5150	101.0240	97.71493